Relationships Between In-Hospital and 30-Day Standardized Hospital Mortality: Implications for Profiling Hospitals

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Objective. To examine the relationship of in-hospital and 30-day mortality rates and the association between in-hospital mortality and hospital discharge practices.

Data Sources/Study Setting. A secondary analysis of data for 13,834 patients with congestive heart failure who were admitted to 30 hospitals in northeast Ohio in 1992–1994.

Design. A retrospective cohort study was conducted.

Data Collection. Demographic and clinical data were collected from patients' medical records and were used to develop multivariable models that estimated the risk of in-hospital and 30-day (post-admission) mortality. Standardized mortality ratios (SMRs) for in-hospital and 30-day mortality were determined by dividing observed death rates by predicted death rates.

Principal Findings. In-hospital SMRs ranged from 0.54 to 1.42, and six hospitals were classified as statistical outliers ($p < .05$); 30-day SMRs ranged from 0.63 to 1.73, and seven hospitals were outliers. Although the correlation between in-hospital SMRs and 30-day SMRs was substantial ($R = 0.78$, $p < .001$), outlier status changed for seven of the 30 hospitals. Nonetheless, changes in outlier status reflected relatively small differences between in-hospital and 30-day SMRs. Rates of discharge to nursing homes or other inpatient facilities varied from 5.4 percent to 34.2 percent across hospitals. However, relationships between discharge rates to such facilities and in-hospital SMRs ($R = 0.08$; $p = .65$) and early post-discharge mortality rates ($R = 0.23$; $p = .21$) were not significant.

Conclusions. SMRs based on in-hospital and 30-day mortality were relatively similar, although classification of hospitals as statistical outliers often differed. However, there was no evidence that in-hospital SMRs were biased by differences in post-discharge mortality or discharge practices.

Key Words. Health policy, health services research, quality of health care, severity of illness index, hospital mortality

Comparisons of hospital mortality rates have generally been based on deaths that occur during hospitalization (i.e., in-hospital deaths) (Hannan, Kilburn, O'Donnell, et al. 1990; Iowa Health Data Commission 1992; Office of Statewide Hospital Planning and Development 1993; Pennsylvania Health Care Containment Council 1991; Rosenthal and Harper 1994) or deaths that occur a standard time (i.e., 30 days) after admission (Sullivan and Toby 1992; Grover et al. 1993). In-hospital deaths can be readily ascertained from patients' medical records or hospital discharge abstracts (i.e., administrative data). However, analyses based only on in-hospital deaths may be subject to bias by differences in length of stay or hospital discharge practices (Kahn, Brook, Draper, et al. 1988). For example, some hospitals may be more adept at discharging patients at high risk of death to nursing homes or hospices. These practices may lead to lower in-hospital mortality rates that are independent of the quality or effectiveness of hospital care.

In contrast, analyses based on deaths that occur over a standard interval after admission may not be subject to such bias. However, ascertaining deaths that occur after discharge is tedious and costly if done through primary data collection. While post-discharge mortality may be available through existing databases, linking information from such databases with hospital information may require the use of unique patient identifiers, such as the social security number, and may raise concerns about risks to patient confidentiality (Donaldson and Lohr 1994; Gostin et al. 1996). Moreover, although some

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databases currently exist that have linked information about hospitalization with vital status information for some patients (e.g., Medicare and Department of Veterans Affairs beneficiaries), these databases are limited to information available on hospital discharge abstracts and may not capture important clinical information that is generally found only in patients' medical records. Comparisons of hospital mortality based on these databases may be subject to bias from unmeasured differences in severity of illness (Green et al. 1990; Iezzoni 1989).

Thus, each of these endpoints for studying hospital mortality has distinct advantages and disadvantages. However, the similarity of hospital comparisons based on in-hospital and post-admission mortality has been poorly studied. Prior studies have been limited by inadequate adjustment for severity of illness (Chassin, Park, Lohr, et al. 1989; Garnick, DeLong, and Luft 1995). In addition, prior studies are based on patients who were hospitalized more than ten years ago and, thus, may not be reflective of current practices that encourage early hospital discharge. Moreover, little is known about the degree to which differences in hospital discharge practices may bias comparisons based on in-hospital mortality.

The current study addressed several unanswered questions about the relationship between in-hospital and 30-day post-admission mortality. We hypothesized that in-hospital and 30-day (post-admission) standardized mortality ratios (SMRs, i.e., observed/predicted mortality) would differ for many of the hospitals in our sample. We further hypothesized that in-hospital SMRs would be biased by differences in discharge practices and by differences in early post-discharge mortality (i.e., deaths occurring after discharge but within 30 days of admission), in ways that would lead to hospitals with lower in-hospital mortality having higher rates of discharge to nursing homes and other facilities and higher rates of early post-discharge mortality.

To determine the predicted mortality rates used in the calculation of hospital SMRs, we used clinical and demographic information abstracted from patients' medical records. These data allowed us to adjust for severity of illness at the time of hospital admission.

METHODS

Patients

The study sample included 13,834 Medicare beneficiaries age 65 years and older with congestive heart failure. Patients were discharged from 30 hospitals in northeast Ohio during the period July 1992 through December 1994. All hospitals were participants in Cleveland Health Quality Choice, a regional initiative to examine hospital performance (Rosenthal and Harper 1994). Five of the participating hospitals were members of the Council of Teaching Hospitals (COTH) of the Association of American Medical Colleges. Other characteristics of participating hospitals have been described previously (Rosenthal and Harper 1994; Rosenthal, Quinn, and Harper 1997).

The sample was drawn in the following manner. First, we identified all 28,196 discharges with congestive heart failure from the Cleveland Health Quality Choice database. This database includes consecutive discharges of patients age 18 years and older with specific ICD-9-CM principal diagnosis codes (398.91, 401.01, 402.11, 402.91, 404.01, 404.03, 404.11, 404.13, 404.91, 404.93, 428.0, 428.1, 428.9), with the exception of patients directly transferred from other acute care hospitals (Rosenthal, Quinn, and Harper 1997). We then merged data for these patients with data from MEDPRO (Medicare Peer Review Organization) files for Medicare beneficiaries age 65 years and older (Ash and Schwartz 1994). The MEDPRO data files include data elements found in the hospital discharge abstract, date of death (as determined from HCFA social security files), and a unique patient identifier: the HCFA identification number (HICN). Because the Cleveland Health Quality Choice database did not include social security numbers or other unique patient identifiers, patients were matched using six elements that were common to both databases, including the dates of hospital admission and discharge, ICD-9-CM principal diagnosis code, gender, discharging hospital, and date of birth (or year of birth). Successful matches were obtained for 21,464 discharges, including 18,275 (85 percent) discharges who were matched on the first five variables and birth date, 1,482 (7 percent) discharges who were matched on the first five variables and birth year, and 1,707 (8 percent) discharges who were matched on the first five variables only.

Upon completing the match, we used the HICN to identify unique combinations of patients and hospitals ($n = 13,834$). For patients with more than one discharge from a particular hospital, we randomly selected a single discharge, and we excluded 7,630 discharges that represented multiple admissions for a single patient to a particular hospital. We chose to randomly select a discharge for patients with multiple discharges because of concerns that selection of an initial admission would automatically select a discharge in which a patient was discharged alive and that selection of the last admission may be biased toward identifying a discharge in which a patient died. Nonetheless, correlations were substantial between hospital mortality rates determined on the basis of selecting initial, random, or last admissions. For example, the correlation between 30-day hospital mortality rates determined using a random admission or the last admission was 0.96, suggesting that differences in criteria for selecting admissions in patients with multiple admissions would not affect study findings.

Data Collection

The Cleveland Health Quality Choice database included information from patients' hospital records that medical records technicians at each hospital abstracted on standard data forms. As previously reported (Rosenthal and Harper 1994; Rosenthal, Quinn, and Harper 1997), several processes were established to ensure data reliability; these included explicit definitions for each variables, double keystroke data entry, identification and correction of variables with missing or out-of-range values, and independent evaluation of the reliability of data abstraction at each hospital. The database included the following data elements: sociodemographics; admission source (e.g., home, nursing home); comorbid conditions; admission medications; admission vital signs and neurological findings; results of laboratory, radiologic, electrocardiographic, and echocardiographic testing from the first 48 hours of admission; dates of "do not resuscitate" orders; and length of stay. Vital status was determined from the Medicare Part A Claims (Modified Medicare Provider Analysis and Review) data files; based on the vital status data, deaths occurring within 30 days after hospital admission were identified, and 30-day post-admission mortality (i.e., 30-day hospital mortality) was determined for each patient.

Analysis

Bivariate associations between demographic and clinical factors and 30-day mortality were examined using the chi-square test for categorical variables and the t-test for continuous and ordinal variables. Factors with significant ($p < .05$) bivariate associations were entered into stepwise logistic regression analyses to identify factors independently related ($p < .01$) to 30-day mortality. In these analyses, continuous variables (e.g., age, systolic blood pressure, serum creatinine) were examined both as continuous data and as interval data. Individual intervals (with the exception of a referent group) were then entered into the logistic regression analyses as indicator (i.e., dummy) variables. Variable classifications that maximized model fit (i.e., discrimination and calibration) were retained in the models. Because our intent was to use the model to adjust mortality rates in the current population and not to export the model for use in other patient populations, we chose not to partition the sample into separate development and validation cohorts and fit models to each cohort. However, given the large sample size of the current study, it is likely that such models would have been similar.

Model discrimination was assessed by the *c*-statistic (Ash and Schwartz 1994), which represents, for all possible pairs of patients who died and who were alive at 30 days, the proportion of times the patient who died had a higher predicted risk of death, based on the logistic regression model. Model calibration was assessed by the Hosmer-Lemeshow statistic, which compares observed and predicted death rates in ten deciles based on the predicted risk of death (Ash and Schwartz 1994). The final risk-adjustment model included 38 variables representing 27 distinct risk factors (Table 2). The relative strengths of the 27 risk factors was assessed by the Wald chi-square statistic (SAS Institute Inc. 1989) For risk factors that were represented by more than one variable, Wald chi-square values for the individual variables were summed.

The risk-adjustment model was used to determine a predicted risk of 30-day death (0 to 100 percent) for each patient. Predicted risks in individual patients were aggregated to determine mean predicted risks for each hospital. Standardized mortality ratios (SMRs) were determined by dividing observed hospital mortality rates by mean predicted mortality rates (Knaus et al. 1993). SMRs greater than 1.0 indicate observed death rates that are higher than expected (i.e., lower hospital performance), whereas SMRs below 1.0 indicate observed death rates that are lower than expected (i.e., higher performance).

In addition, we estimated a predicted risk of in-hospital death for each patient by fitting a second logistic regression model to the same variables in the 30-day mortality model, and determined in-hospital SMRs. Confidence intervals around 30-day and in-hospital SMRs were estimated by calculating exact 95 percent limits around observed mortality rates and dividing these by the mean predicted mortality rate, which was taken as a constant. Hospitals were classified as outliers if the 95 percent confidence interval did not include 1.0.

Associations between 30-day and in-hospital SMRs and between in-hospital SMRs and early post-discharge mortality rates were examined using the Spearman correlation coefficient; results based on the Pearson correlation coefficient were nearly identical. In addition, to examine the potential effect of differences in discharge practices on hospital mortality, hospitals were grouped into quintiles on the basis of in-hospital SMRs. Mean rates of early post-discharge mortality and mean rates of discharge to nursing homes or other inpatient facilities (acute care hospitals, skilled nursing facilities, hospices, and rehabilitation hospitals) were then compared across quintiles using analysis of variance (ANOVA). Analyses were conducted using SAS for Windows, Version 6.12.

RESULTS

The mean age of the 13,834 hospital admissions was 79 years. Fifty-eight percent of persons admitted were female and 83 percent were white (Table 1). The most common comorbid conditions included ischemic heart disease in 56 percent of patients, diabetes mellitus in 32 percent, and chronic obstructive lung disease in 27 percent. Seventy-three percent of patients were discharged home, 20 percent were discharged to skilled nursing facilities or to other acute care hospitals, and 7.5 percent of patients died in the hospital. Of the 12,795 patients discharged alive, 99 percent (n = 12,679) were discharged within 30 days of admission, and of the 1,039 patients who died in the hospital, 97 percent (n = 1,005) died within 30 days of hospital admission.

The 30-day mortality rate was 12.0 percent. Of the 1,655 patients who died within 30 days of admission, 63 percent died during the initial hospital admission. Of the 12,679 patients who were discharged from the hospital within 30 days of admission, 5.1 percent (n = 650) died after discharge and within 30 days of the index admission (i.e., early post-discharge mortality). Early post-discharge mortality rates were lower in patients discharged home than in patients discharged to nursing homes or other inpatient facilities (3.0 percent versus 11.1 percent, respectively, p = .001). Sixty-day and 90-day mortality rates were 17.8 percent and 21.5 percent, respectively.

Predictors of 30-Day Hospital Mortality

Bivariate and multivariate odds ratios of the variables included in the logistic regression model for 30-day mortality are shown in Table 2. Of the 27 risk factors included in the model, the ten that were the strongest multivariable predictors included the systolic blood pressure at admission (Wald chi-square, 240.1), age (72.9), lowest serum sodium during the first 48 hours (53.9), highest blood urea nitrogen during the first 48 hours (47.2), cancer (47.0), admission neurological assessment (46.6), serum aspartate aminotransferase (30.4), serum albumin (28.5), admission from a nursing home (25.4), and the most abnormal arterial pCO_2 during the first 48 hours (24.9). The c-statistic of the logistic regression model was 0.795; the Hosmer-Lemeshow statistic

Table 1: Sociodemographic and Clinical Characteristics of 13,834 Study Patients

Mean Age (years) ± s.d.	78.9 ± 7.7	
Median (interquartile range)	78 (73 – 84)	
Mean Predicted Risk of Death ± s.d. (30 day)	0.12 ± 0.13	
Median (interquartile range)	0.07 (0.04 – 0.14)	
Mean Predicted Risk of Death ± s.d. (In-Hospital)	0.08 ± 0.11	
Median (interquartile range)	0.04 (0.03 – 0.06)	
Mean Hospital Length of Stay (days) ± s.d.	7.5 ± 6.4	
Median (interquartile range)	6 (4 – 9)	
	Percent of Patients	*(Number)*
Gender		
Male	41.7	(5,766)
Female	58.3	(8,068)
Race		
White	83.2	(11,503)
African American	15.7	(2,165)
Other/Not documented*	1.2	(162)
Part B Medicare Coverage	67.5	(9,339)
Admission Source		
Home	85.9	(11,872)
Skilled nursing facility	13.3	(1,836)
Other/Not documented	0.8	(107)
Admission from Emergency Room	73.5	(10,163)
Comorbid Conditions		
Prior stroke/TIA	17.6	(2,436)
Ischemic heart disease	56.1	(7,766)
Diabetes mellitus	32.4	(4,487)
Chronic obstructive lung disease	27.3	(3,780)
Cancer (metastatic or receiving chemotherapy)	3.0	(412)
Cirrhosis	0.5	(71)
End-Stage renal disease (chronic dialysis)	1.2	(172)
Level of Consciousness on Admission		
Alert	94.0	(13,001)
Lethargy or stupor	3.5	(485)
Coma	0.8	(107)
Not documented	1.7	(234)
Hospital Discharge Disposition		
Home	48.4	(6,696)
Home with home health care	23.8	(3,295)
Skilled nursing or rehabilitation facility	17.4	(2,402)
Transfer to acute care hospital	2.4	(325)
Discharge against medical advice	0.3	(43)
Discharged alive but location not documented	0.2	(34)
In-Hospital death	7.5	(1,039)
In-Hospital death within 30 days of admission	7.3	(1,005)

for this model was not significant (chi-square = 13.8, 8 df, $p = .08$), indicating that the model was well calibrated (Table 3). The *c*-statistic of a second logistic regression estimating the risk of in-hospital mortality, based on the same 27 risk factors, was 0.818.

Relationship Between In-Hospital and 30-Day Mortality

Thirty-day mortality rates exhibited more than threefold variation across hospitals, ranging from 6.4 percent to 20.4 percent; mean predicted 30-day mortality rates, based on the logistic regression model, ranged from 9.0 percent to 15.0 percent; 30-day SMRs ranged from 0.63 to 1.73. Seven hospitals were classified as outliers—that is, SMRs were higher or lower ($p < .05$) than 1.0—including four high and three low outliers. In-hospital mortality rates also varied more than threefold, ranging from 3.2 percent to 10.7 percent. In-hospital SMRs ranged from 0.54 to 1.42; six hospitals were classified as outliers, including two low and four high outliers. Although the correlation between 30-day SMRs and in-hospital SMRs was substantial ($R = 0.78$, $P < .001$; Figure 1), classification of hospitals as outliers often differed. Only three hospitals (all high outliers) were outliers on the basis of both 30-day and in-hospital SMRs. Outlier status changed for seven hospitals (23 percent), including four hospitals classified as outliers on the basis of 30-days SMRs (one high and three low outliers) that were non-outliers on the basis of in-hospital SMRs, and three hospitals that were non-outliers on the basis of 30-day SMRs that were outliers on the basis of in-hospital SMRs (one high and two low outliers). However, absolute values of the difference between 30-day and in-hospital SMRs were not higher in the seven hospitals in which outlier status changed than were the corresponding absolute values in the 23 hospitals in which outlier status did not change (mean values, 0.12 versus 0.13, respectively, $p = .85$). Of the seven hospitals in which outlier status changed, differences between in-hospital and 30-day SMRs were less than 0.10 in three hospitals, and were 0.10 to 0.20 in three hospitals; in only one hospital was the difference greater than 0.20. The above results were nearly identical in analyses that excluded the 1,836 patients who were admitted from nursing homes.

Influence of Early Post-Discharge Mortality and Discharge Practices

Rates of early post-discharge mortality varied more than fivefold, ranging from 2.4 percent to 12.9 percent. However, rates of early post-discharge

Table 2: Bivariate and Multivariate Associations with 30-Day Mortality of Variables Included in the Multivariable Rick-Adjustment Models

			Bivariate Associations		Multivariable Associations	
Variable		*Number of Patients*	*30-Day Mortality Rate (Percent)*	p-*Value*	*Odds Ratio (95% CI)*	p-*Value*
Demographic Characteristics						
Age (years)	65–69	2,022	7.6	<.001	—*	—
	70–74	2,766	9.6		1.23 (0.98–1.54)	.07
	75–79	3,093	10.1		1.32 (1.06–1.64)	.01
	80–84	2,798	13.2		1.68 (1.36–2.09)	<.001
	85–89	1,886	16.1		2.00 (1.60–2.53)	<.001
	≥ 90	1,269	19.8		2.52 (1.97–3.21)	<.001
Admission from nursing home	*Yes*	1,836	24.9	<.001	1.47 (1.26–1.70)	<.001
	No	11,998	10.0		—*	
Comorbid Conditions						
Cancer (metastatic or receiving chemotherapy)	*Yes*	412	25.7	<.001	2.48 (1.91–3.21)	<.001
	No	13,422	11.5		—*	
Chronic obstructive lung disease/bronchitis	*Yes*	3,780	13.5	<.001	1.20 (1.06–1.37)	.005
	No	10,054	11.4		—*	
Prior stroke or transient ischemic attack	*Yes*	2,436	14.4	<.001	1.14 (0.99–1.32)	.08
	No	11,398	11.4		—*	
Cirrhosis	*Yes*	71	21.1	.017	1.99 (1.06–3.75)	.03
	No	13,763	11.9		—*	
Use of systemic corticosteroids prior to admission	*Yes*	771	16.3	<.001	1.24 (0.99–1.56)	.06
	No	13,063	11.7		—*	

Vital Signs/Conditions Present at Admission						
Cardiac arrest at time of hospital admission	*Yes*	56	39.3	.001	1.92 (1.00–3.69)	.05
	No	13,778	11.9		—*	
Mechanical ventilation on day of admission	*Yes*	485	23.7	<.001	1.37 (1.03–1.83)	.03
	No	13,349	11.5		—*	
Admission neurological assessment	*Coma*	107	55.1	<.001	3.01 (1.90–4.80)	<.001
	Lethargy/stupor	485	34.2		1.79 (1.42–2.25)	<.001
	Alert	13,242	10.8		—*	
Admission heart rate < 40 beats/minute	*Yes*	27	22.2	.036	2.46 (0.89–6.78)	.08
	No	13,807	11.9		—*	
Admission respirations ≥ 30 beats/minute	*Yes*	3,384	14.5	<.001	1.21 (1.06–1.39)	.005
	No	10,450	11.1		—*	
Admission systolic blood pressure (mm Hg)	≤ 79	49	55.1	<.001	0.98 (0.98–0.99) (per 1 mm Hg increase)	<.001
	80–99	429	35.2			
	100–119	1,813	20.6			
	120–149	4,942	12.5			
	≥ 150	6,601	7.3			
Admission Chest X-Ray Findings						
Presence of infiltrate	*Yes*	3,150	16.2	<.001	1.22 (1.08–1.39)	.002
	No	10,684	10.7		—*	
Admission Laboratory Findings						
Serum creatinine (mg/dl)	< 1.5	8,071	8.8	<.001	1.23 (1.07–1.42) (per increase in category)	.004
	1.5–2.4	4,091	14.6			
	≥ 2.5	1,672	20.9			

continued

Table 2: Continued

Variable		*Number of Patients*	*Bivariate Associations* 30-Day Mortality Rate (Percent)	*Bivariate Associations* p-Value	*Multivariable Associations* Odds Ratio (95% CI)	*Multivariable Associations* p-Value
Serum blood urea nitrogen (mg/dl)	< 20	4,198	5.8	<.001	—*	—
	20–29	4,408	8.7		1.20 (0.99–1.46)	.06
	30–39	2,225	14.1		1.56 (1.23–1.98)	<.001
	40–49	1,113	17.5		1.63 (1.21–2.20)	.001
	≥ 50	1,890	27.5		2.19 (1.55–3.08)	<.001
Blood urea nitrogen/ creatinine ratio	< 15	3,947	7.7	<.001	1.11 (1.05–1.19) (per increase in category)	.001
	15–19	4,215	8.7			
	20–24	3,039	12.1			
	25–29	1,332	19.1			
	≥ 30	1,301	27.7			
Lowest serum sodium (meq/l)	≥ 135	11,090	10.4	<.001	—*	—
	130–134	1,930	16.4		1.28 (1.10–1.48)	.001
	120–129	741	21.7		1.73 (1.40–2.13)	<.001
	≤ 119	73	30.1		3.25 (1.87–5.67)	<.001
Highest serum sodium ≥ 150 meq/l	*Yes*	140	37.9	<.001	1.70 (1.11–2.61)	.01
	No	13,694	11.7		—*	
Serum lactate dehydrogenase (LDH; IU/l)	≤ 249	9,673	9.7	<.001	1.22 (1.08–1.37) (per increase in category)	.001
	250–999	3,914	16.4			
	≥ 1,000	247	31.2			
Serum aspartate amino transferase (SGOT; IU/l)	≤ 39	10,644	9.7	<.001	1.22 (1.14–1.31) (per increase in category)	<.001
	40–79	2,347	15.6			
	80–119	363	24.2			
	120–399	346	28.6			
	≥ 400	134	47.8			

White blood count (X 1,000)	$\leq$ 12	10,755	9.8	<.001	1.15 (1.08–1.23) (per increase in category)	<.001
	12–15	1,611	16.8			
	15–20	967	20.8			
	20–30	410	24.4			
	$\geq$ 30	91	31.9			
Serum albumin (mg/dl)	$\leq$ 2.4	420	26.7	<.001	0.73 (0.66–0.82) (per 1 mg/dl increase in serum albumin)	<.001
	2.5–2.9	1,327	19.6			
	3.0–3.4	3,500	12.8			
	3.5–3.9	6,522	10.8			
	$\geq$ 4.0	2,065	6.3			
Serum bicarbonate $\geq$ 40 meq/l	*Yes*	70	24.3	<.001	1.84 (1.00–3.38)	.05
	No	13,764	11.9		—*	
Arterial pH < 7.30	*Yes*	979	21.5	<.001	1.34 (1.07–1.67)	.01
	No	12,855	11.2		—*	
Arterial pCO_2 (mm Hg)	< 25 or $\geq$ 60	764	24.3	<.001	1.58 (1.27–1.97)	<.001
	25–29 or 50–59	1,536	17.6		1.27 (1.08–1.50)	.004
	30–50	11,534	11.2		—*	
Arterial pO_2 < 60 (mm Hg)	*Yes*	2,666	14.9	<.001	1.20 (1.05–1.38)	.008
	No	11,168	11.3		—*	

*Referent category for variables expressed using one or more dummy (i.e., indicator) variables.

Table 3: Calibration of the Multivariable Model Used to Estimate the Risk of 30-Day Mortality

		Dead (N=*1,655*)		*Alive* (N=*12,179*)	
Severity Decile	*Number of Patients*	*Observed Rate (N)*	*Predicted Rate (N)*	*Observed Rate (N)*	*Predicted Rate (N)*
1	1386	0.019 (27)	0.017 (23.6)	0.983 (1359)	0.981 (1362.4)
2	1384	0.020 (27)	0.028 (39.2)	0.972 (1357)	0.980 (1344.8)
3	1383	0.034 (47)	0.038 (53.1)	0.962 (1336)	0.966 (1329.9)
4	1375	0.042 (58)	0.049 (68.0)	0.951 (1317)	0.958 (1307.0)
5	1384	0.056 (77)	0.063 (87.3)	0.937 (1307)	0.944 (1296.7)
6	1384	0.087 (120)	0.081 (112.1)	0.919 (1264)	0.913 (1271.9)
7	1381	0.117 (162)	0.107 (147.4)	0.893 (1219)	0.883 (1233.6)
8	1385	0.162 (224)	0.144 (199.7)	0.856 (1161)	0.838 (1185.3)
9	1383	0.215 (298)	0.217 (300.7)	0.783 (1085)	0.785 (1082.3)
10	1389	0.443 (615)	0.449 (623.3)	0.551 (774)	0.557 (765.7)

Note: The overall Hosmer-Lemeshow statistic (see Methods) was not significant (chi-square = 13.9; 8 df; $p = .08$), indicating that the model was well calibrated.

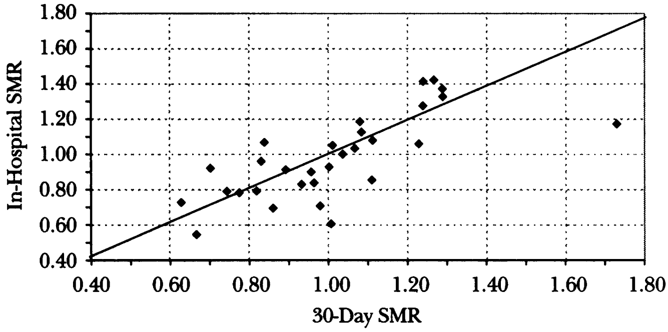
mortality were not related to in-hospital SMRs ($R = 0.18$; $p = .34$). When hospitals were further stratified into quintiles on the basis of in-hospital SMRs, mean rates of early post-discharge mortality in the first through fifth quintiles (i.e., first quintile includes hospitals with the six lowest SMRs) were 5.2 percent, 5.4 percent, 4.7 percent, 6.4 percent, and 5.5 percent, respectively ($p = .72$, ANOVA). Thus, no evidence was present to suggest that hospitals with lower in-hospital SMRs had higher early post-discharge mortality.

The proportion of patients discharged to nursing homes or other inpatient facilities also varied widely, ranging from 5.4 percent to 34.2 percent. Proportions were lower than 10 percent in two hospitals and higher than 25 percent in six hospitals. However, proportions were not related to either in-hospital SMRs ($R = 0.08$; $p = .65$) or early post-discharge mortality ($R = 0.23$; $p = .21$). Finally, when hospitals were stratified into quintiles on the basis of in-hospital SMRs, mean proportions of patients discharged to nursing homes or other inpatient facilities in the first through fifth quintiles were similar (21.7 percent, 20.0 percent, 20.2 percent, 21.7 percent, and 16.1 percent, respectively; $p = .57$, ANOVA).

DISCUSSION

Recent efforts to decrease hospital utilization have led to marked declines in length of stay and to increases in the use of skilled nursing facilities and

Figure 1: Relationship Between 30-Day and In-Hospital Standardized Mortality Ratios (SMRs).



Note: For each hospital, SMRs were determined by dividing observed 30-day or in-hospital mortality by the mean predicted 30-day or in-hospital mortality. Predicted mortality rates were determined from multivariable models (see Methods). The correlation between SMRs was significant (Spearman $R = 0.78$; $p < .001$). The solid diagonal line represents equal in-hospital and 30-day SMRs.

other postacute settings for delivery of care (Prospective Payment Assessment Commission 1995). Such changes in hospital utilization have important implications for the use of in-hospital mortality as an indicator of hospital performance or quality of care. (Hannan, Kilburn, O'Donnell, et al. 1990; Iowa Health Data Commission 1992; Office of Statewide Hospital Planning and Development 1993; Pennsylvania Health Care Containment Council 1991; Rosenthal and Harper 1994). Indeed, such profiles may be biased by differences in the ways in which hospitals utilize nursing homes, hospices, or other postacute care settings for patients with poor short-term prognoses. The current study examined the potential importance of these factors by comparing in-hospital and 30-day SMRs for patients with congestive heart failure from 30 hospitals. We emphasize the following findings.

First, both in-hospital and 30-day SMRs exhibited nearly threefold variation across hospitals. In addition, substantial variation in discharge practices was observed; rates of discharge to nursing homes and other inpatient facilities varied more than sixfold across the 30 hospitals. Second, although most deaths within 30 days of admission occurred in the hospital, nearly 40 percent of the deaths occurred after hospital discharge. However, in contrast to our hypothesized relationship, hospitals with lower in-hospital SMRs did not have higher post-discharge mortality rates, nor were hospitals with lower in-hospital mortality rates more likely to discharge patients to nursing homes or other inpatient facilities. Moreover, in-hospital and 30-day SMRs were strongly correlated, again in contrast to our hypothesized relationship. Nonetheless, the classification of hospitals as statistical outliers on the basis of in-hospital and 30-day SMRs differed for seven of the 30 hospitals, although absolute differences in SMRs in these seven hospitals were relatively small, suggesting that such hospitals had SMRs that were close to the limits of the 95 percent confidence intervals used to determine outlier status.

Taken together, the findings of this study suggest that in-hospital mortality may still represent a reasonably valid marker for 30-day mortality and may be relatively unbiased despite differences in hospital discharge practices. These findings are also consistent with findings from two earlier studies based on Medicare claims data. Chassin, Park, Lohr, et al. (1989) found that in-hospital and 30-day mortality rates were strongly correlated for patients discharged in 1984 with 48 medical and surgical conditions, although mortality rates were adjusted for age, race, and gender only. For all conditions, the correlation between hospital rankings for the two endpoints was 0.54, while for congestive heart failure, the correlation was 0.71. In addition, the authors found that nearly half of the hospitals that were in the lower fifth percentile for in-hospital mortality had a similar classification based on 30-day mortality. Garnick, DeLong, and Luft (1995) examined HCFA hospital mortality data for patients discharged in 1989 with congestive heart failure and found that 30-day and 180-day hospital mortality were strongly correlated ($R = 0.74$), but that 30-day mortality and mortality that occurred 31 to 180 days after admission were only weakly correlated. Although mortality rates were adjusted for severity using administrative data only, and although this study did not examine in-hospital mortality, the findings for 30-day and 180-day mortality rates mimic the current findings for in-hospital and 30-day mortality rates.

Thus, the current findings add to the two earlier studies in important ways. First, the study involved a more contemporary cohort of patients and may be more reflective of current patterns of hospital utilization. Second, the current study adjusted for admission severity of illness using more sophisticated risk-adjustment models that were developed from clinical data abstracted from patients' medical records. Finally, the current study specifically examined the potential impact of variations in discharge practices.

However, the current findings may be inconsistent with earlier studies of the effect of differences in hospital discharge practices. Jencks, Williams, and Kay (1988) found that in-hospital mortality for Medicare patients hospitalized in 1985 with stroke, pneumonia, acute myocardial infarction, and congestive heart failure was 25 percent higher in New York than in California, but that length of stay was almost twice as long in New York as in California and that 30-day mortality rates were actually similar in the two states. Kahn, Keeler, Sherwood, et al. (1990) studied Medicare patients hospitalized before (1981/1982) and after (1985/1986) the implementation of prospective payment and found that, although severity-adjusted in-hospital mortality declined nearly 20 percent, length of stay declined nearly 25 percent and 30-day post-admission mortality declined by only 6 percent.

Although the studies by Jencks, Williams, and Kay (1988) and Kahn, Keeler, Sherwood, et al. (1990) suggest that in-hospital mortality may be biased by differences in discharge practices, these studies involved patients who were hospitalized more than ten years ago. Moreover, the length of stay of patients included in the earlier studies and the variation in length of stay (e.g., nearly 50 percent shorter in California than New York) were much greater than in the current study. In addition, these studies did not examine the effect of differences in discharge practices at a hospital-specific level and determined neither the relationship between in-hospital and 30-day SMRs nor the classification of hospitals as statistical outliers.

In interpreting our findings, several potential methodological limitations should be considered. First, although to adjust for severity of illness we used multivariable models based on clinical data that were abstracted from patients' medical records and that had excellent discrimination, relative to current methods (Iezzoni 1989; Ash and Schwartz 1994; Steen, Brewster, Bradbury et al. 1993), it is possible that unmeasured severity varied across hospitals and/or across the two endpoints. However, such differences would likely decrease associations between in-hospital and 30-day mortality. Thus, our results actually may underestimate the true association between the two endpoints. Second, because we used clinically based severity models, our findings may represent a "best case" scenario. It is possible that associations between in-hospital and 30-day mortality SMRs based on other severity measures that explain a lower proportion of the variation in mortality—such as measures that utilize administrative data—would be lower. Third, our analyses were based on admissions that occurred over a 30-month period. Such long sample frames yield relatively large hospital volumes that may mitigate the effect of random variation in mortality. It is possible that associations between in-hospital and 30-day mortality would have been lower if hospital volumes had been smaller and the effect of random variation had been larger (Park, Brook, Kosecoff, et al. 1990; Rosenthal et al. 1998). Fourth, the generalizability of our findings to regions that have different patterns of hospital utilization and managed care penetration should be established. In addition, although the 30 hospitals in our sample included a diverse spectrum of hospitals in terms of teaching status, size, and other characteristics, the generalizability of our study to regions with different types of hospitals should also be examined. Finally, it is important that future investigations examine the applicability of our findings to other diagnoses, particularly to surgical conditions and conditions associated with either higher or lower short-term mortality.

In spite of the above potential limitations, the current findings may have important implications for the use of mortality data to profile hospital performance. If generalizable to other conditions and geographic regions, the findings suggest that standardized mortality rates based on in-hospital and 30-day mortality may be relatively similar, despite differences in early post-discharge mortality and differences in rates of discharge to nursing homes and other facilities. Nonetheless, it is important to recognize that statistical classifications of hospitals as high- or low-mortality outliers, as is commonly done in reporting hospital performance (Hannan, Kilburn, O'Donnell, et al. 1990; Iowa Health Data Commission 1992; Office of Statewide Hospital Planning and Development 1993; Pennsylvania Health Care Containment Council 1991; Rosenthal and Harper 1994; Sullivan and Toby 1992; Grover et al. 1993), may differ, and that such differences may have substantial implications for public perceptions of hospital quality and for hospital contracting by employers and managed care. However, the differences in outlier status may also result from random variations in mortality rates (Park, Brook, Kosecoff, et al. 1990; Rosenthal et al. 1998) or from differences in the width of statistical confidence intervals that result from the greater number of deaths included in 30-day SMRs than in in-hospital SMRs.

Thus, the current findings have important implications for initiatives that use mortality rates as a measure of hospital performance, particularly initiatives that rely on data abstracted from patients' medical records and that do not have ready access to patients' vital status after discharge. The collection of vital status following hospital discharge may entail considerable resources and effort, and is best done using unique patient identifiers. However, as noted earlier, the inclusion of patient identifiers in healthcare databases that contain sensitive clinical information raises concerns about patient confidentiality (Donaldson and Lohr 1994; Gostin et al. 1996). No national consensus currently exists on an optimal identifier for inclusion in longitudinal healthcare databases, and, although identifiers, such as the social security number, are often routinely collected by health insurance companies and governmental agencies, a recent Institute of Medicine Task Force did not recommend their use (Donaldson and Lohr 1994). In addition, although post-discharge vital status data are available for certain patients, such as Medicare beneficiaries for whom hospital claims can be linked to social security data files, such data are less readily available for patients with commercial health insurance or Medicaid, or for those without insurance.

In the absence of easily accessible healthcare databases with longitudinal patient data, the examination of hospital-based endpoints may represent a cost-effective means for measuring hospital performance. However, the potential limitations of such hospital-based endpoints must always be considered. Moreover, as the care of patients is further shifted away from acute care settings and to skilled nursing facilities and patients' homes, it is important that these issues be examined again in the future.

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